WPS5846

Policy Research Working Paper 5846

Tall Claims

Mortality Selection and the Height of Children

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October 2011

Abstract

Data from three rounds of nationally representative health surveys in India are used to assess the impact of selective mortality on children's anthropometrics. The nutritional status of the child population was simulated under the counterfactual scenario that all children who died in the first three years of life were alive at the time of measurement. The simulations demonstrate that the difference in anthropometrics due to selective mortality would be large only if there were very large differences in anthropometrics between the children who died and those who survived. Differences of this size are not substantiated by the research on the degree of association between mortality and malnutrition. The study shows that although mortality risk is higher among malnourished children, selective mortality has only a minor impact on the measured nutritional status of children or on that status distinguished by gender.

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TALL CLAIMS:
MORTALITY SELECTION AND THE HEIGHT OF CHILDREN

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Keywords: Mortality, nutrition, children, India

JEL: J12, I32, O12

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I. Introduction

Anthropometric status is often used as an indicator of welfare (Steckel 1995, 2009). Many people also consider adequate nutrition to be of intrinsic value—a good in its own right. Others have emphasized its instrumental role in productivity and economic growth (Fogel 2004; Alderman, Behrman, and Hoddinott 2005). But regardless of whether nutritional status is used to explain welfare or to understand growth, it may be important to determine whether trends or differences in anthropometric measures of nutritional status in population groups[1] are a reflection of trends and differences in another key indicator of welfare, mortality.

For example, Deaton (2007) explores the relationship of income and adult height and speculates on explanations for a positive relationship between mortality and height in Africa. This relationship, which contrasts with that found for all other countries pooled in a separate regression, is observed even after accounting for country specific effects. One possible reason for this association is that in Africa mortality selection dominates the what Deaton refers to as "scarring" – a reduction in adult height because of disease and malnutrition in childhood, Mortality selection is, of course, an extreme form of sample attrition, widely recognized as a potential source of bias in economic analysis (Fitzgerald, Gottschalk, and Moffitt 1998). The well documented association of malnutrition and mortality (see, for example, Victora and others 2008) may truncate the lower tail of the height distribution, and this may be sufficient to offset the negative association of child mortality with the causes of malnutrition.

While Deaton's study is an attempt to explain the pattern of heights in Africa, the concern that selection may mask patterns of health is more general. For example, in a review dominated by evidence from developed countries Almond and Currie (2009), write: "*Finally, Bozzoli, Deaton, and Quintana-Domeque [2009] highlight that in developing countries, high average mortality rates cause the selection effect of early childhood mortality to overwhelm the 'scarring' effect. Thus, the positive relationship between early childhood health and subsequent human capital may be absent in analyses that do not account for selective attrition in high mortality settings*". Almond and Currie, however, do not offer functional definition for 'high'. Results of studies on nutrition may be called into question when mortality is high. For example, Maccini and Yang (2009) worry about the possible bias in their estimates of heights that might come from selective mortality (though they offer a simple argument that this is not a concern for Indonesia). Our paper is partially motivated by the view that the legitimate concern for extreme mortality environments might be taken out of their range of validity.

[1] While anthropometry does not cover all aspects of nutrition – many micronutrient deficiencies do not manifest in changes in weight or height – it is a commonly tracked measure. Unless otherwise stated, this paper implies anthropometric measures of nutritional status when discussing nutrition and malnutrition.

Steckel (2009) suggests that this impact of mortality on the height of survivors can be tested through simulations. The current paper is in keeping with this strategy. Although it does not simulate the effect of sample truncation on nutritional status across countries, it does employ simulations at the individual level that can indicate whether selective mortality might mask improvements in nutritional status over time in a country where mortality rates have been declining rapidly or whether truncation might bias comparisons across genders. Three rounds of nationally representative health survey data from India are used to simulate the nutritional status of the child population under the polar counterfactual that all children who died in the first three years of life are alive at the time of measurement.

The objective of this study is similar to that of studies by Boerma et al (1992), Pitt (1997) and Dancer, Rammohan, and Smith (2008), although the approach differs. Boerma et al. use data from longitudinal studies and retrospective data from cross-sectional studies in 17 countries to analyze the effect of selective survival on children's anthropometric measures. Their study concludes that selective survival has only a marginal effect on the comparisons of anthropometric outcomes across geographic areas, subpopulation and time. Pitt recognizes that children who fail to survive through childhood are not a random draw from a population and furthermore that fertility itself is a choice. He addresses this by simultaneously estimating the probability of these two events along with nutritional status and finds that although fertility and mortality are statistically significant determinates of nutrition there is no behaviorally significant bias in the parameters if this selection is ignored. Dancer, Rammohan, and Smith use a selection correction to estimate models of nutritional status and find that survival is positively associated with nutritional status. That is, they find that scarring, to use Deaton's terminology, is more prevalent in the sample than is selection.

Similarly, a paper by Gorgens et al. (2007) finds evidence that extremely high mortality rates during the 1959-196 famine in China impacts trends in adult height in rural areas, but the affect of the estimated selection is relatively small. At the same time, no significant mortality bias was found for urban population in China. Bozzoli, Deaton and Quintana-Domeque (2009) also demonstrate that the selection could have a significant positive impact on adult height at very high mortality rates[2]. A recent paper by Moradi (2010) estimates the size of the selection affect of survival in Gambia and finds it to be too small to account for the tall adult heights observed in Sub-Saharan Africa.

[2] The paper presents several specifications of the regression of adult height on pre-adult mortality rates and mortality rates squared. The coefficients on the linear mortality rates are negative and those on quadratic mortality rates are positive and significant. The estimated inflection rate after which higher pre-adult mortality rates have a positive effect on adult height varies by specification, but for all specifications this rate is outside of the data range (i.e. at or above pre-adult 250 deaths per 1000).

The current study confirms that even when mortality risk is higher among malnourished children, this has only a minor impact on the measured nutritional status of the child population or on that status by gender with evidence from a country with high rates of malnutrition and moderately high mortality. We show this, first, by illustrating the degree to which the nutrition results reported in various surveys form India would have changed had mortality rates differed. While our initial simulations impute results using global evidence on the relationship of nutrition and risk of mortality we also bolster these simulations with additional simulations based on the estimated hazard of mortality from survey data.

The paper is organized as follows: The next section describes the data and presents some descriptive statistics. Section II outlines the theoretical framework and the empirical strategy. Section III discusses the main results, and section IV presents some implications of the findings.

II. Data and Descriptive Statistics

This analysis uses data from three waves of India's National Family Health Survey (NFHS; 1992/93, 1998/99 and 2005/06), a survey of representative households in states and territories covering some 99 percent of the population[3] and similar in structure to demographic and health surveys conducted in several other countries. The NFHS follows the pattern of a standard Demographic and Health Survey. The main sample of NFHS contains information on 45,279 children in 33,032 households from the 1992/93 round, 30,984 children in 26,056 household from the 1998/99 round, and 48,679 children in 33,968 households from the 2005/06 round.[4]

Because the NFHS does not collect information on household income or consumption, a household wealth index was constructed from the data on household assets using the method based on principal components (see, for example, Filmer and Pritchett 2001; Rutstein and Johnson 2004).

The NFHS provides height and weight data for children under age 48 months in 1992/93, under age 36 months in 1998/99, and under age 60 months in 2005/06. The NFHS contains no anthropometric information for deceased children at the time of their death. For comparability between NFHS rounds, the sample was restricted to children under age 36 months. The analysis focuses on the age-adjusted measure of height-for-

[3] Kashmir, Sikkim, and some remote territories were not covered in NFHS-1. Detailed information on NFHS methodology and sample design is available at www.nfhsindia.org/. The data are available from MEASURE DHS, Macro International Inc. at www.measuredhs.com.

[4] The number of observations in the 1998/99 round is smaller, as it collected height and weight information only for the last two children under age 3 of ever-married women who were interviewed. In the 1992/93 round, measurements of height were not collected in Andhra Pradesh, Himachal Pradesh, Madhya Pradesh, Tamil Nadu, and West Bengal.

age, which reflects children's development relative to a reference population of well-nourished children (WHO 2006).[5]

Because children's weight at birth influences their health and prospects for survival (Rosenzweig and Shultz 1982), a key variable for this study is the data on children's weight at birth in the survey. The NFHS collects information on weight at birth in addition to weight at the time of the survey and asks mothers to categorize the weight of their children at birth as large, average, or small.[6] The sample of children with weight measured at birth is much smaller than the sample of children whose weight was assessed by their mother (table A1 in the appendix). The study relies mainly on these subjective assessments as a proxy for health endowments at birth. The weight of about 20–25 percent of children was lower than average according to their mothers' assessments (see table A1), a figure not out of keeping with the rate of low birth weight children in India.

The data show that average height-for-age has risen over time for both boys and girls, with z-scores rising from –1.91 in 1992 to –1.55 in 2005 for boys and from –1.86 to –1.53 for girls. Correspondingly, stunting declined from 72 percent of boys and 70 percent of girls in 1995 to 65 percent for both sexes by 2005. Despite these improvements, malnutrition remains prevalent in India.

Deaths around the time of birth (neonatal) are high in all rounds of the NFHS (table 1). Boys accounted for about 56 percent of all deaths in the immediate postnatal period. However, beyond age 6 months, the share of deaths is higher for girls than for boys. A plot of the cumulative mortality hazard by gender for the three rounds of the NFHS again shows that more than half the deaths in children under age 36 months occur in the first month after birth, with little change over the years. It also shows a higher mortality risk for boys than for girls in the early months of life (figure 1), followed by a reversal in later months. This switch in mortality patterns results in almost identical total mortality rates for boys and girls ages 0–36 months.

[5] A report by Nutrition Foundation of India concluded that the World Health Organization (WHO) standard was generally applicable to Indian children (IIPS 2000). The nutritional status of children calculated in this way is compared with the nutritional status of an international reference population recommended by the WHO (Dibley and others 1987). The use of this reference group is based on the empirical finding that well-nourished children in all population groups for which data exist follow similar growth patterns (Martorell and Habicht 1986). Across rounds of the NFHS, about 10 percent of eligible children were not measured, either because the children were not at home or because their mothers refused to allow the measurements (Lokshin, Das Gupta, and Ivaschenko 2005).

[6] There is a good correspondence between the measured weight and the weight at birth assessed by the mothers. For example, in 1992, only about 3 percent of children who were assessed as large at birth had measured weight in the lowest quintile of the weight distribution. Seidman and others (1987) show that about 75 percent of self-reported birth weights were accurate within 100 grams. A study by Adegboyea and Heitmannb (2008) that uses a weight categorization similar to NFHS concludes that maternal assessment of a child's weight at birth "seems to be sufficiently accurate for clinical and epidemiological use."

III. Quantifying the Impact of Selective Mortality on Nutritional Indicators

The magnitude of the influence of selective mortality on nutritional status of a population depends on child mortality level, prevalence of malnutrition among living children, and prevalence of malnutrition among deceased children (Boerma et al. 1992). Our population – all children – consists of living and deceased children. Taking a height-for-age *z*-score as a measure of malnutrition, the average height-for-age *z*-score for all children Z_{all} can be expressed as:

(1) $$Z_{all} = Z_s(1 - P_d) + Z_d P_d,$$

where Z_s and Z_d are average *z*-scores for survivors and deceased children, correspondingly, and P_d is the proportion of dead children. The change in the *z*-score had deceased children survived is then:

(2) $$\Delta = Z_s - Z_{all} = (Z_s - Z_d)P_d$$

Assuming that at least some deaths are related to malnutrition ($Z_d < Z_s$) the inclusion of dead children in the sample results in lower average z-scores ($Z_{all} < Z_s$). Δ is larger the higher is child mortality and the larger is the difference in the average z-scores of living and deceased children.

The analysis starts with simple simulations illustrating empirically the magnitude of the potential influence of selective mortality and then moves on to simulations based on a proportional hazard model.

Simulations Illustrating the Magnitude of Potential Impact of Selective Mortality

What would be the observed height-for-age for all Indian children ages 0–36 months had the children who died before age 36 months survived? Table 2 presents the simulated changes in the average height-for-age z-score for different imputation scenarios for three rounds of NFHS.

The first set of results demonstrates how large the difference in z-scores between children who died before age 36 months (Z_d) and those who survived (Z_s) should for Δ in (2) to be statistically significant. In 1992, 7.7 percent of boys and 8.0 percent of girls died before age 36 months. The average height-for-age z-score of surviving children (Z_s) was –1.91 with a standard error of 0.016 for boys and –1.86 with a standard error of 0.017 for girls. The first row of table 2 shows, that imputing a z-score of -2.5 for dead boys results in the statistically significant changes in overall z-score (Z_{all}) from the actual -1.91 to -1.95. For girls, the impact of selective mortality would be statistically significant had the currently deceased girls survived and their average z-score were -2.3.

If the average height-for-age of children who died was twice as far below the age and gender reference mean as that of children who survived (–3.83 rather than –1.91), the height-for-age z-score for the total sample would rise from –1.91(Z_s) to –2.17(Z_{all}), or by 13.5 percent. If a height-for-age z-score of –5.0 (the lower bound recommended as the cut-off for outliers; WHO 1995) is imputed to a sample of children who died, the overall z-score would rise from –1.91 to –2.32, or by 21.8 percent. Similar tendencies are observed for the 1998 and 2005 samples.

The impact of the imputations for height-for-age on the total sample is proportional to the mortality rates. For example, the imputation of a height-for-age z-score twice as low as the average to the sample of girls who died before age 36 months results in a 13.0 percent change in the overall mean z-score in 1992 but only a 8.4 percent change in 1998 and a 6.6 percent change in 2005, reflecting the decline in girls' mortality rates from 0.077 in 1992 to 0.056 in 2005.

But not all deaths before age 36 months were caused by malnutrition. The next simulation is based on the results from the literature that estimates the contribution of malnutrition to child mortality. Puffer and Serrano (1973) found that malnutrition was an underlying cause in 54 percent of deaths for children ages 2–4 years. Pelletier (1994) explored 28 prospective datasets and found that the population-attributable risk of mortality associated with anthropometric deficits varied from 17 percent to 74 percent in eight studies for Asia and Africa. Pelletier and others (1994) applied to prospective surveys in Ethiopia, Guatemala, India, and Malawi a new methodology for determining the association of malnutrition and mortality by the severity of malnutrition. They demonstrated that 42–57 percent of deaths of children ages 6–59 months were associated with malnutrition's potentiating effects on infectious disease, 76–89 percent of them attributable to mild to moderate malnutrition. Analysis of data for 53 developing countries for the 1980s found that about 56 percent of child deaths were associated with malnutrition. The proportion is close to 67 percent for India, with 73–74 percent of it attributable to mild to moderate malnutrition (Pelletier and others 1995).

While this approach is based on weight for age and not height-for-age, it makes a good starting point for a simulation of imputed height-for-age, based on the assumption that 67 percent of deaths in children up to age 36 months in India were related to malnutrition (the upper bound for that association; see bottom panel of table 2). The height-for-age of living children by gender in a particular year was used to impute the average height-for-age of children whose deaths were not associated with malnutrition. Of deaths among children related to malnutrition, 70 percent were assumed to have occurred among children with moderate to mild malnutrition, and an average height to age z-score of –2.5 was imputed to them. For the rest of the sample of children who had died, an average height to age z-score of –4 was imputed[7].

Even though these imputations are based on the upper bound estimates of Pelletier and others (1995), they result in only modest changes in the overall mean. The largest changes are observed for the 1992 sample because mortality rates are highest in that year (see bottom panel of table 2). Had all the children who died survived, the total mean height-for-age z-score for boys would change from –1.91 to –2.01, a 4.9 percent change. The impact of imputations is smaller for 1998, at 3.6 percent and 2005, at 4.1 percent. For girls, the imputations change the total mean from –1.86 to –1.95 in 1992, a 5.1 percent change, and by 3.3 percent for 1998 and 4.09 percent for 2005. The change in height-for-age z-score is slightly larger in percentage terms in 2005 than in 1998 (although the absolute value of the change is smaller) in keeping with the lower mortality rate.

Recent studies by Pelletier and Frongillo (2003) and Black and other (2008) indicate that globally among children younger than 36 months the proportion of deaths associated with malnutrition declined to 37 percent in the late 1990s and early 2000s because of the effect of expanded coverage of immunizations, oral rehydration therapy, antibiotics, and other child survival interventions. Imputations of height-for-age z-scores corrected for contemporaneous mortality selection based on these estimates would yield smaller and statistically insignificant changes in the total mean for the 1998 and 2005 NFHS samples. While these more modest associations are not illustrated in table 2, the simulations that are shown demonstrate that the selectivity mortality would only have a large impact on observed anthropometrics if there were very large differences in anthropometrics between the children who died and those who survived. Current research on the association between mortality and malnutrition does not substantiate such large differences.

Simulations Based on a Proportional Hazard Model

The simulations discussed in the previous section were based on imputations of a few categories of height-for-age data to the children in the sample who had died. It is more realistic to assume that children's anthropometrics and survival depend on their individual characteristics, prenatal conditions, health of the mother, and those of their household (Wolpin 1997). To approximate this, children who died were matched with children who survived past the age of 36 months using the estimated survival hazard as a matching score. The anthropometric scores of children who survived were then imputed

[7] Boerma et al (1992) demonstrate that the differences in malnutrition between the children who died after the measurement and who survived a specified time period in longitudinal studies in Indian states of Tamil Nadu and Punjab were relatively small: height-for-age z-scores of 60 percent of dead children were lower than -2 SD from the mean compared to 50 percent among survivors.

to the children who had died, and the impact of these imputations on the average height-for-age z-score of the total sample was estimated.

The estimation uses a standard theoretical framework of household utility maximization that incorporates the production function of a child's health (Behrman and Deolalikar 1988). Household utility is a function of the consumption and leisure of household members and the quality (health) and quantity of their children. A household maximizes its utility subject to budget constraints and the restrictions imposed by the health production function. The household demand for child health at time *t* depends on a set of exogenous characteristics of the child and its mother, household, and community, as well as some unobserved factors captured by the random error term ε_{it} (Thomas, Strauss, and Henriques 1991). This relation can be expressed, in linear form, as:

$$H_{it} = \beta \bar{X}_{it} + \varepsilon_{it} \tag{3}$$

where vector $\bar{X}_i$ combines the child's, mother's, household's, and community's characteristics, and β is a vector of parameters.

The child's health can be linked to mortality through the stochastic rule for observing death (Sickles and Taubman 1997). The mortality state for child *i* at time *t* is defined as:

$$M_{it} \begin{cases} = 1 \;\; if \;\; H_{it} < H_{it}^{*} \\ = 0 \;\; otherwise \end{cases} \tag{4}$$

where H_{it}^{*} is a child- and time-specific mortality threshold that can be interpreted as a shock whose arrival time follows a Poisson distribution. The probability that a shock H_{it}^{*} occurs during the period $(t, t+\Delta)$ is $P = h_0(t)\Delta + o(\Delta)$. Then the hazard of dying during this period is:

$$h_i(t) = h_0(t)[1 - F(H_{it})] \tag{5}$$

where $h_0(t)$ is the baseline hazard and $F(H_{it})$ is the health distribution function. The survival function is then:

$$S_i(t \mid x_i) = \exp\{-h_0(t)[1 - F(H_{it})]t\}. \tag{6}$$

Assumptions about the distribution of the health shocks that are standard in the literature on mortality can be used to estimate the survival function using the Weibull proportion hazard model, such that:

$$\begin{aligned} & h_i(t) = \theta t^{\theta-1} \exp\{-H_{it}\} = \theta t^{\theta-1} \exp\{-\beta x_i\} \\ & and \\ & S(t \mid x_i) = \exp\{-\theta t^{\theta-1} \exp(-\beta x_i) t\} \end{aligned} \tag{7}$$

where θ is an unknown parameter.

To control for the unobserved heterogeneity in children's health (frailty), a parameter is introduced that represents the effect of unobserved factors on survival. This effect is not directly estimated from data but is assumed to have unit mean and variance α_i that is estimated (see Vaupel, Manton, and Stallard 1979; Lancaster 1979; Hougaard 1995). In the presence of such unobserved heterogeneity, survival function (7) becomes:

(8) $$S(t \mid x_i, \alpha_i) = \{S(t \mid x_i)\}^{\alpha_i} .$$

The unconditional survival function is obtained by integrating out the unobservable factor by assuming a distribution for α. If the unobserved factor is inverse-Gaussian-distributed,[8] the survival function of the Weibull proportional hazard model (8) becomes:

(9) $$S(t \mid x_i, \alpha) = \exp\{\frac{1-[1-2\alpha \ln(S(t))]^{1/2}}{\alpha}\}.$$

The parameters of equation (9) are estimated using the maximum likelihood algorithm. Once these parameters are obtained, dead and surviving children can be matched on the estimated hazard as a matching score, and this matching can be used to impute a height-for-age z-score for children who died. The matching approach used here is analogous to the approach used in propensity score matching (see Rubin 1973, 1979; Rosenbaum and Rubin 1983), but the hazard function is a different functional form than commonly used for propensity score matching. It was chosen for its ability to account for censoring and truncation, making it the preferred model for estimating survival.[9] The overall conclusion is robust to an estimation of survival based on a probit regression (results not reported in tables), the functional form more commonly employed in propensity score matching.

IV. Results

The explanatory variables used in proportional hazard estimation (equation 7) include the child's sex, birth order, and weight at birth; the mother's age, educational attainment, employment status, and other characteristics; household size, socio-demographic

8 For the inverse-Gaussian frailty distribution, the relative variability of frailties among surviving children decreases with age, which could be a more realistic assumption for modeling child mortality; Gamma frailty distribution assumes constant variability of frailty with age (Gutierrez 2002).

9 Samples of dead and surviving children were matched using a nearest neighbor algorithm with the restriction that observations in both samples are on a common support in terms of the matching score (Heckman, LaLonde, and Smith 1999). The probability density functions of matching scores for dead and surviving children are shown in figure A1 in the appendix. An alternative, and probably more intuitive, way to run these simulations would be to model children's height for age z-scores as a function of their characteristics and to impute height for age z-scores to the dead children using out-of-sample prediction. The approach here is similar to that because the out-of-sample prediction could be interpreted as a case of matching. The advantages of the current approach are the use of a more flexible function form for the matching and the ability of the hazard model to deal better with attrition issues.

composition, wealth index, religion, and caste; and the availability of community services and infrastructure. The descriptive statistics for these variables are presented in table 3.

Tables 4 and 5 show the coefficients of the Weibull proportional hazard estimation for children younger than age 36 months using three rounds of the NFHS.[10] Table 4 includes the child's weight at birth estimated by the mother; Table 5 uses weight measured at birth. The weight at birth is one of the important factors affecting neo-natal child mortality that combines the unobserved information about pre-natal conditions and shocks experienced by the mother and her child (i.e., Claeson, et al. 2000). The results of the two estimations are not directly comparable as one is a categorical variable (mother's assessment of weight at birth) and the other is a continuous variable (measured weight), but the implications are similar[11]. The remainder of the discussion focuses on the estimations based on the specification using mother-assessed weight at birth because the sample sizes are much larger.

The estimated hazard odds ratios on the control variables in the estimations of proportional hazard models reveal the expected relationship between child mortality and characteristics of the child, mother, and household. A child's gender has no effect on mortality hazard: mortality rates for boys are higher than for girls in the six months after birth and lower after that. A higher birth order has a negative impact on survival probabilities (Miller et al. 1992). Weight at birth, whether assessed by the mother or measured at birth, is a strong predictor of mortality. Children whose mother's assessed their weight as small (with an odds ratio greater than 1 on the dummy variable reported in table 4) and children with low measured weight (with odds ratios less than one on the continuous measured weight variables in table 6) are significantly less likely to survive than are children who weigh more at birth.

Children living in the wealthiest households and with better educated mothers have better prospects for survival than do children from poor households and with less educated mothers. In an inverted U-shaped relationship, children's survival improves with the mother's age till about age 40 and declines thereafter.

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[10] The specification with Weibull distribution is selected based on the comparison of Akaike (1974) information criterion values for specifications with exponential, Weibull, Gompertz, log-normal, log-logistic, and general gamma distributions.

[11] Note that the interpretation of the numbers presented in tables 5 and 6 are different from the standard interpretation of the regression coefficients. The odds ratios are always positive; odds ratios that are less than one indicate that an increase in a particular factor reduces the probability of an event, while odds ratios greater than 1 mean that a particular factor increases the probability of an event. Correspondingly, the t-tests of the odds ratio test the null hypothesis of odds ratios being equal to 1. For comparability, we re-estimated the hazard model shown in Table 5 with the continuous variable on weight at birth categorized as a binary indicator equal to 1 for children with weight at birth < 2,500 grams. The odds ratio for this dummy is less than 1 and significant, but the effect of being born with low weight is smaller compared to coefficient in Table 5. One explanation for this could be that children whose weight was measured at birth come from the wealthier families with better access to post-natal health care.

Table 6 presents the simulated impact of the matched imputations of height-for-age z-scores for all children ages 0–36 month who did not survive till the age of 36 months. The average z-score is higher for children with measured weight at birth than for children with mother-assessed weights at birth reflecting wealth differences.

In 1992, the average height-for-age z-score for boys who survived past age 36 months was –1.913. When the matched z-score estimates are imputed to children who died before age 36 months, the overall average height-for-age z-score becomes –1.936; the difference of –0.023 is not statistically significantly different from 0. Similar differences are observed for other years. In all years, for both boys and girls, the imputations have no significant impact on the overall anthropometric indices. In all cases, the average imputed height-for-age z-score in the simulations based on the survival model is smaller than the imputed z-score that would result in a statistically significant change in the overall height-for-age z-score as shown in first panel of table 2[12]. These results resonate with other studies that use longitudinal data from surviving and non-surviving children and find a modest amount of selection via child mortality on height-for-age z-scores i.e., Boerma et al (1992), Moradi (2010)[13].

The effect of the change in mortality on gender patterns of height-for-age z-scores was also simulated using the results of hazard function estimations. This simulation closes the gender gap in mortality in the age 3–36 month group by artificially increasing the mortality among boys[14]. The simulation assumes that 144 of the boys in the sample with the lowest probabilities of survival did not survive and thus did not contribute to the observed height-for-age z-scores. The simulation of this increase in boys' mortality on the data from 1992 of NFHS demonstrates a 0.12 percent increase in the height-for-age z-score. This clearly has a negligible impact on the difference in nutritional status between boys and girls.

Finally, taking our model a step further, we simulate the health outcomes of Indian children if the mortality rates in India were as high as one of the highest rates current in Africa. In this simulation we change the surviving status of living children with lowest probabilities of survival to reach the mortality rate of 15 percent for both boys and

[12] Table A2 in Appendix presents the simulated impact of the matched imputations of height-for-age z-scores for children ages 3–36 months who did not survive till the age of 36 months. Neonatal mortality was excluded in the hazard estimates as it could have a different set of correlates, but the results are not particularly sensitive to the exclusion of this age group.

[13] Our empirical model fails to account for the possible selection bias due to high levels of maternal mortality in India (e.g., Ronsmans and Graham 2006). Correction of this bias would increase the gap in anthropometric outcomes between the deceased and living children. Unfortunately, the nature of IFHS sample makes it difficult to look on the relationship of maternal mortality and children health outcomes.

[14] In principle, one could, simulate a more desirable decrease in girls' mortality to close this gap, but it is more direct to use the estimates from boys actually in the sample then to make projections for girls who would otherwise have been in the sample.

girls.[15] The results of the simulation (table 7) show only minor changes in the aggregate height-for-age z-scores when the mortality rates are increased from 5 – 8 percent (table 1) to 15 percent. The changes are statistically significant for the samples of boys and girls in 1998 and for the sample of boys in 2005, but even in these cases the magnitudes of the changes have not exceeded 3 percent.

V. Conclusion

This paper used data from three rounds of a nationally representative health survey in India to assess the magnitude of the bias in children's anthropometrics due to selective mortality. The nutritional status of the child population was simulated under the counterfactual scenario that all children who had died in the first three years of life were alive at the time of measurement. These simulations imputed various values of height-for-age z-scores to the sample of dead children. The simple simulations, with imputed z-scores that are independent of the child's characteristics, show that, at the rates of child mortality prevailing in India in 1992–2005, the selective mortality could have only a moderate impact on overall anthropometric measures. The imputations based on the literature on the association between child mortality and nutrition result in only a 5 percent difference between the counterfactual and the actual height-for-age z-scores.

The simulations based on the hazard model that takes into account differences in mortality and anthropometrics related to child characteristics are consistent with the observation that malnourished children are less likely to survive and thus to contribute to anthropomorphic measurements. However, the results also show that with the current low (and declining) mortality rates by historical standards, improved survival rates have an insignificant impact on overall height-for-age z-scores. The *changes* in mortality between 1992 and 2005 imply that some malnourished children who would previously have died instead survived and are measured in the 2005 survey. To the degree that selective mortality affects overall malnutrition levels or rates, the reductions in child mortality mask some of the improvement in nutrition. However, the results of this study suggest that progress on the fourth Millennium Development Goal to reduce child mortality only lightly obscures the results for the target for the first Millennium Development Goal to reduce malnutrition and hunger.[16]

Similarly, the findings imply that differences in mortality are unlikely to explain gender differences in anthropometrics—or their absence. While the NFHS data do not

[15] This rate corresponds to the high mortality rate of 14.2 percent observed among boys in Cote d'Ivoire in 1998. The rate for Cote D'Ivoire was the highest mortality rate from 30 DHS of African countries reviewed.

[16] This target is measured in terms of underweight rather than height for age. Underweight is somewhat easier to measure in a survey covering young children, though height for age is a clearer indicator of cumulative health. While underweight rates and stunting rates often differ, trends in the two tend to move together.

show a marked gender pattern in overall mortality of children 0–36 months, deaths are higher for girls after the neonatal period. Nevertheless, the imputations here have no significant impact on relative nutritional status. While the results reported here are from India over a two decade period, in a more general sense, it appears that selective mortality is unlikely to be of significant magnitude in most countries to have a large impact on trends in populations or subpopulations.

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Table 1: Proportion of total deaths by age and gender for three rounds of India's National Family Health Survey (percent)

	1992		1998		2005	
Age	**Boys**	**Girls**	**Boys**	**Girls**	**Boys**	**Girls**
Neonatal	50.7	40.6	53.75	44.0	56.9	47.3
0–6 months	19.7	20.7	18.9	19.4	18.6	19.6
7–12 months	14.2	17. 8	12. 5	16.2	10.9	14.5
13–18 months	2.2	2.7	2.0	2.6	1.7	2.3
19–24 months	8.4	11.6	8.2	11.5	7.4	10.6
25–36 months	4.9	6.6	4.7	6.3	4.5	5.6
Total	100	100	100	100	100	100
Mortality rate	0.080	0.077	0.064	0.066	0.058	0.056
Standard error	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)

Source: Authors' analysis based on data from India's National Family Health Survey.

Table 2: **Changes in the total mean height-for-age z-score (HAZ) for different imputation scenarios.**

	Boys			Girls		
	Imputed HAZ	Mean HAZ Actual → Simulated	%Δ total HAZ	Imputed HAZ	Mean HAZ Actual → Simulated	%Δ total HAZ
	Z_d	$Z_s \rightarrow Z_{all}$		Z_d	$Z_s \rightarrow Z_{all}$	
Imputed HAZ (Z_d) that results in a statistically significant change in total HAZ (Z_{all})						
1992	-2.15	-1.91 → -1.95	1.70	-2.13	-1.86 → -1.90	1.85
1998	-2.12	-1.77 → -1.80	1.58	-2.22	-1.85 → -1.88	1.65
2005	-1.95	-1.55 → -1.57	1.77	-1.98	-1.53 → -1.56	1.92
Imputed HAZ (Z_d) is twice as low as the average observed HAZ (Z_s)						
1992	-3.83	-1.91 → -2.17	13.51	-3.72	-1.86 → -2.10	12.99
1998	-3.55	-1.77 → -1.92	8.12	-3.71	-1.85 → -2.01	8.37
2005	-3.09	-1.55 → -1.65	6.78	-3.06	-1.53 → -1.63	6.60
Imputed HAZ (Z_d) = - 5SD						
1992	-5.00	-1.91 → -2.33	21.81	-5.00	-1.86 → -2.27	21.90
1998	-5.00	-1.77 → -2.03	14.78	-5.00	-1.85 → -2.12	14.21
2005	-5.00	-1.55 → -1.78	15.13	-5.00	-1.53 → -1.76	14.95
HAZ imputed based on upper bounds of Pelletier et al. (1995) estimates						
1992	-2.61	-1.91 → -2.01	4.91	-2.59	-1.86 → -1.95	5.09
1998	-2.56	-1.77 → -1.84	3.61	-2.59	-1.85 → -1.92	3.32
2005	-2.49	-1.55 → -1.61	4.12	-2.48	-1.53 → -1.59	4.09

Source: Authors' analysis based on data from India's National Family Health Survey and estimates from Pelletier and others (1995) on the upper bound of deaths in children due to malnutrition.

Table 3: Descriptive statistics for the main explanatory variables.

Variable	1992 Mean	1992 Standard error	1998 Mean	1998 Standard error	2005 Mean	2005 Standard error
Male child	0.509	*0.003*	0.519	*0.003*	0.523	*0.003*
Child's current age (in months)	17.326	*0.054*	17.556	*0.058*	18.062	*0.061*
Birth order						
First	0.275	*0.002*	0.292	*0.002*	0.311	*0.003*
Second	0.239	*0.002*	0.259	*0.002*	0.276	*0.003*
Third	0.177	*0.002*	0.176	*0.002*	0.160	*0.002*
Fourth	0.116	*0.002*	0.105	*0.002*	0.096	*0.002*
Fifth	0.075	*0.001*	0.069	*0.001*	0.064	*0.001*
Sixth	0.048	*0.001*	0.044	*0.001*	0.039	*0.001*
Seventh	0.031	*0.001*	0.025	*0.001*	0.022	*0.001*
Eighth	0.038	*0.001*	0.031	*0.001*	0.031	*0.001*
Mother's current age (in years)	25.878	*0.029*	25.512	*0.029*	25.831	*0.030*
Education of the mother (years)	2.703	*0.022*	3.506	*0.025*	4.243	*0.028*
Education of the mother (category)						
No education	0.641	*0.002*	0.540	*0.003*	0.487	*0.003*
Incomplete primary	0.144	*0.002*	0.090	*0.002*	0.065	*0.001*
Complete primary	0.037	*0.001*	0.069	*0.001*	0.071	*0.001*
Incomplete secondary	0.129	*0.002*	0.159	*0.002*	0.284	*0.003*
Complete secondary	0.020	*0.001*	0.063	*0.001*	0.040	*0.001*
Higher	0.028	*0.001*	0.080	*0.001*	0.053	*0.001*
Scheduled caste	0.133	*0.002*	0.202	*0.002*	0.212	*0.002*
Scheduled tribe	0.093	*0.001*	0.096	*0.002*	0.098	*0.002*
Religion						
Hindu	0.795	*0.002*	0.793	*0.002*	0.802	*0.002*
Muslim	0.153	*0.002*	0.158	*0.002*	0.153	*0.002*
Christian	0.020	*0.001*	0.023	*0.001*	0.019	*0.001*
Sikh	0.018	*0.001*	0.014	*0.001*	0.013	*0.001*
Other or no religion	0.015	*0.001*	0.012	*0.001*	0.014	*0.001*
Wealth index score	–0.306	*0.004*	–0.293	*0.005*	-0.451	*0.005*
Urban	0.226	*0.002*	0.223	*0.002*	0.255	*0.003*
Household size	7.931	*0.021*	7.682	*0.022*	7.049	*0.020*
Share of children ages 0–6	0.314	*0.001*	0.318	*0.001*	0.329	*0.001*
Share of children ages 7–14	0.131	*0.001*	0.121	*0.001*	0.116	*0.001*
Share of men	0.246	*0.001*	0.250	*0.001*	0.238	*0.001*
Share of women	0.246	*0.001*	0.250	*0.001*	0.238	*0.001*
Share of elderly	0.035	*0.000*	0.034	*0.000*	0.033	*0.000*
Type of toilet						
Flush	0.164	*0.002*	0.186	*0.002*	0.308	*0.003*
Latrine	0.081	*0.001*	0.112	*0.002*	0.041	*0.001*
Other or none	0.755	*0.002*	0.701	*0.002*	0.651	*0.003*
Source of drinking water						
Piped	0.279	*0.002*	0.321	*0.003*	0.321	*0.003*
Well or hand pump	0.665	*0.002*	0.645	*0.003*	0.643	*0.003*
Surface, river, rain, other	0.038	*0.001*	0.027	*0.001*	0.027	*0.001*
Number of observations	37,558		33,547		29,798	

Note: Sample includes children younger than age 36 months with nonmissing height-for-age data. Other explanatory variables include 27 state dummy variables.
Source: Authors' analysis based on data from India's National Family Health Survey.

Table 4: Weibull proportional hazard estimation for children ages 0–36 months using mother's estimate of birth size.

Variable	1992 Hazard ratio	1992 Standard error	1998 Hazard ratio	1998 Standard error	2005 Hazard ratio	2005 Standard error
Male child	1.050	*0.049*	1.002	*0.054*	0.987	*0.062*
Birth order						
Second	1.724***	*0.124*	2.289***	*0.234*	2.053***	*0.192*
Third	2.442***	*0.233*	3.638***	*0.470*	2.579***	*0.344*
Fourth	3.472***	*0.393*	5.135***	*0.832*	3.338***	*0.568*
Fifth	3.768***	*0.517*	5.437***	*0.994*	5.102***	*0.918*
Sixth	3.968***	*0.646*	6.309***	*1.347*	6.937***	*1.456*
Seventh	4.337***	*0.765*	7.648***	*1.788*	7.502***	*1.920*
Eighth	7.718***	*1.369*	11.271***	*2.682*	9.086***	*2.430*
Born Small	2.008***	*0.109*	1.493***	*0.088*	1.381***	*0.100*
Characteristics of the mother						
Mother's age (in years)	0.941*	*0.031*	0.915**	*0.035*	1.045	*0.057*
Mother's age squared	1.001	*0.001*	1.000	*0.001*	0.998**	*0.001*
Incomplete primary	0.780***	*0.061*	1.098	*0.103*	1.034	*0.128*
Complete primary	0.867	*0.119*	0.747**	*0.092*	0.824	*0.101*
Incomplete secondary	0.770***	*0.074*	0.825**	*0.079*	0.642***	*0.061*
Complete secondary	0.502**	*0.141*	0.710**	*0.106*	0.607**	*0.138*
University and higher	0.496***	*0.118*	0.591***	*0.115*	0.325***	*0.080*
Household characteristics						
Household size	0.772***	*0.013*	0.756***	*0.016*	0.694***	*0.017*
Household size squared	1.007***	*0.001*	1.007***	*0.001*	1.011***	*0.001*
Share of children ages 0– 6	0.001***	*0.000*	0.001***	*0.000*	0.001***	*0.000*
Share of children ages 7–15	0.177***	*0.051*	0.107***	*0.036*	0.182***	*0.071*
Share of elderly (60+)	0.303***	*0.111*	0.140***	*0.062*	0.435*	*0.215*
Share of women	0.552*	*0.172*	0.418***	*0.137*	0.544	*0.218*
Second wealth quintiles	1.053	*0.068*	0.926	*0.068*	1.148	*0.104*
Third wealth quintiles	1.021	*0.076*	0.838**	*0.070*	1.121	*0.118*
Fourth wealth quintiles	0.921	*0.080*	0.848	*0.094*	1.003	*0.135*
Fifth wealth quintiles	0.634***	*0.097*	0.639***	*0.105*	0.983	*0.184*
Hindu religion	0.944	*0.151*	0.931	*0.148*	1.007	*0.194*
Muslim religion	1.069	*0.182*	0.961	*0.171*	1.072	*0.225*
Scheduled caste	1.231***	*0.081*	1.104	*0.075*	0.994	*0.079*
Scheduled tribe	0.915	*0.077*	0.892	*0.083*	0.845	*0.100*
Living conditions						
Latrine	0.765*	*0.105*	0.912	*0.119*	1.209	*0.218*
Other/none	0.908	*0.114*	0.911	*0.108*	1.146	*0.124*
Well/handpump	0.911	*0.063*	1.083	*0.080*	1.076	*0.107*
Surface, river, rain	1.007	*0.126*	0.998	*0.161*	1.529**	*0.277*
Other	0.764	*0.137*	0.757	*0.244*	1.283	*0.409*
Urban	0.929	*0.075*	0.979	*0.092*	1.017	*0.095*
Constant	1.408	*0.906*	4.352**	*2.727*	0.400	*0.360*
Ln(p)	0.533***	*0.008*	0.530***	*0.010*	0.481***	*0.007*
Ln(θ)	0.155	*0.176*	0.454	*0.278*	0.340	*0.190*
Log-Likelihood		–13,870.68		–10,181.48		–9,308.23
Number of observations		37,621		33,556		29,859

*** p<0.01, ** p<0.05, * p<0.1.

Note: Coefficients on dummy variables for 27 states are not shown.

Source: Authors' analysis based on data from India's National Family Health Survey.

Table 5: Weibull proportional hazard estimation for children ages 0–36 months using recorded weights.

Variable	1992 Hazard ratio	1992 Standard error	1998 Hazard ratio	1998 Standard error	2005 Hazard ratio	2005 Standard error
Male child	1.588**	*0.287*	1.181	*0.186*	1.133	*0.168*
Birth order						
Second	1.967***	*0.479*	4.476***	*0.986*	2.983***	*0.547*
Third	3.994***	*1.287*	4.205***	*1.334*	4.367***	*1.212*
Fourth	7.157***	*3.013*	4.286***	*2.062*	9.430***	*3.451*
Fifth	11.530***	*6.806*	10.650***	*5.736*	3.081*	*1.946*
Sixth	0.000***	*0.000*	5.201**	*3.669*	49.997***	*26.524*
Seventh	1.277	*1.473*	17.193***	*11.702*	30.526***	*16.715*
Eighth	16.028***	*13.796*	10.887**	*12.939*	1.214	*1.539*
Measured weight at birth	0.457***	*0.068*	0.533***	*0.074*	0.668***	*0.098*
Characteristics of the mother						
Mother's age (in years)	0.905	*0.107*	0.905	*0.103*	1.293*	*0.193*
Mother's age squared	1.001	*0.002*	1.001	*0.002*	0.994**	*0.003*
Incomplete primary	0.930	*0.238*	0.821	*0.190*	1.047	*0.257*
Complete primary	1.358	*0.426*	0.533*	*0.196*	0.763	*0.247*
Incomplete secondary	0.642	*0.182*	0.461***	*0.105*	0.555***	*0.104*
Complete secondary	0.791	*0.324*	0.477**	*0.141*	0.469**	*0.147*
University and higher	0.512	*0.210*	0.317***	*0.100*	0.209***	*0.078*
Household characteristics						
Household size	0.756***	*0.046*	0.668***	*0.038*	0.684***	*0.042*
Household size squared	1.006***	*0.001*	1.012***	*0.002*	1.010***	*0.003*
Share of children ages 0–6 years	0.001***	*0.000*	0.001***	*0.000*	0.001***	*0.000*
Share of children ages 7–15	0.063***	*0.063*	0.075***	*0.070*	0.022***	*0.018*
Share of elderly (60+)	0.637	*0.741*	0.078**	*0.092*	0.222*	*0.194*
Share of women	0.467	*0.424*	0.581	*0.484*	0.214**	*0.150*
Second wealth quintiles	0.694	*0.320*	0.787	*0.250*	1.147	*0.327*
Third wealth quintiles	0.653	*0.289*	0.884	*0.271*	1.628*	*0.426*
Fourth wealth quintiles	0.613	*0.276*	0.872	*0.279*	1.915**	*0.542*
Fifth wealth quintiles	0.399*	*0.201*	1.235	*0.454*	2.078**	*0.686*
Hindu religion	0.835	*0.221*	1.194	*0.348*	0.846	*0.234*
Muslim religion	0.670	*0.271*	1.011	*0.395*	0.802	*0.287*
Scheduled caste	0.663	*0.247*	1.038	*0.218*	1.084	*0.191*
Scheduled tribe	0.426*	*0.190*	1.011	*0.326*	0.756	*0.232*
Living conditions						
Latrine	0.819	*0.243*	0.981	*0.280*	0.957	*0.358*
Other/none	1.228	*0.325*	0.928	*0.249*	1.344	*0.279*
Well/handpump	0.938	*0.177*	0.916	*0.158*	1.351**	*0.194*
Surface, river, rain	1.301	*0.686*	1.154	*0.697*	1.043	*0.559*
Other	0.394	*0.238*	2.168*	*0.990*	1.342	*0.837*
Urban	1.546*	*0.346*	0.888	*0.213*	1.071	*0.184*
Constant	22.884*	*39.678*	55.530**	*100.703*	0.339	*0.744*
Ln(p)	0.547***	*0.027*	0.517***	*0.023*	0.522***	*0.018*
Ln(θ)	0.001***	*0.000*	0.001***	*0.000*	0.001***	*0.000*
Log-Likelihood		–1,011.30		–1,239.06		–1,791.04
Number of observations		6,228		8,555		12,755

***p<0.01, **p<0.05, *p<0.1.

Note: Coefficients on dummy variables for 27 states are not shown.

Source: Authors' analysis based on data from India's National Family Health Survey

Table 6: The simulated impact of the matched imputations. Difference between actual and simulated height-for-age z-scores by gender for three rounds of India's National Family Health Survey, children ages 0–36 months

	Boys				*Girls*			
	Imputed HAZ	Mean HAZ Actual → Simulated	Δ	*Std. Error*	Imputed HAZ	Mean HAZ Actual → Simulated	Δ	*Std. Error*
Sample of children with weight at birth assessed by their mothers								
1992	-2.083	-1.913 → -1.936	0.023	0.023	-2.046	-1.863 → -1.887	0.024	0.024
1998	-1.880	-1.771 → -1.779	0.009	0.020	-2.017	-1.848 → -1.862	0.014	0.022
2005	-1.568	-1.554 → -1.555	0.001	0.020	-1.665	-1.537 → -1.545	0.09	0.021
Sample of children with measured weight at birth								
1992	-1.435	-1.284 → -1.295	0.011	0.048	-1.220	-1.282 → -1.278	-0.003	0.050
1998	-1.541	-1.291 → -1.301	0.010	0.033	-1.384	-1.279 → -1.282	0.003	0.037
2005	-1.549	-1.199 → -1.213	0.014	0.029	-1.205	-1.186 → -1.186	0.001	0.031

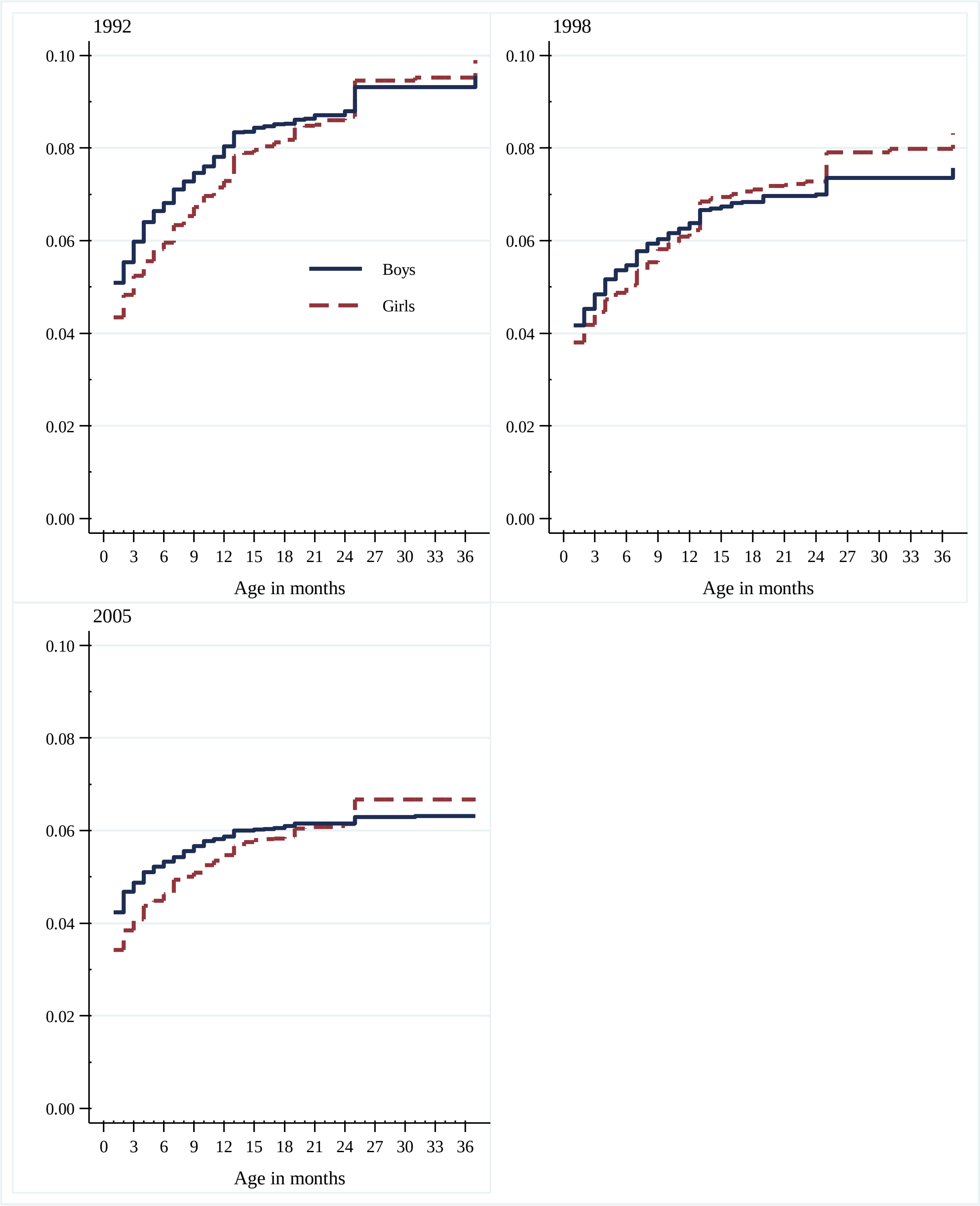
Note: The standard errors for the differences are not adjusted for the fact that propensity score is estimated.
Source: Authors' analysis based on data from India's National Family Health Survey.

Table 7: The simulated child mortality rate equals to high mortality rates in Africa (15 percent). Difference between actual and simulated height-for-age z-scores by gender for three rounds of India's National Family Health Survey, children ages 0–36 months

	Boys			*Girls*		
	Mean HAZ Actual → Simulated	Δ	*p-value*	Mean HAZ Actual → Simulated	Δ	*p-value*
Sample of children with weight at birth assessed by their mothers						
1992	-1.913 → -1.884	-0.029	0.103	-1.863 → -1.840	0.022	0.205
1998	-1.771 → -1.723	-0.047*	0.010	-1.848 → -1.820	0.028*	0.077
2005	-1.554 → -1.533	-0.020	0.129	-1.537 → -1.517	0.019	0.178
Sample of children with measured weight at birth						
1992	-1.284 → -1.241	-0.042	0.289	-1.282 → -1.243	*-0.039*	0.330
1998	-1.291 → -1.260	-0.032	0.190	-1.279 → -1.237	0.041	0.119
2005	-1.199 → -1.162	-0.036*	0.092	-1.186 → -1.161	0.025	0.286

Note: The p-values are for the test on the statistical differences between the actual and simulated mean; *p<0.1
Source: Authors' analysis based on data from India's National Family Health Survey.

Figure 1: Nelson-Aalen cumulative hazard by gender for three rounds of India's National Family Health Survey



Source: Authors' analysis based on data from India's National Family Health Survey.

Appendix

Table A1: Distribution of weight at birth for children ages 0–36 months at time of India National Family Health Surveys by gender, mothers-assessed weight and measured weight

	Boys			Girls		
Year and mothers-assessed weight	**Share**	**Measured weight Mean (kg)**	**Standard error**	**Share**	**Measured weight Mean (kg)**	**Standard error**
1992						
Larger than average	0.143	3.334	0.026	0.126	3.252	0.031
Average	0.649	2.855	0.013	0.636	2.747	0.013
Smaller than average	0.208	2.131	0.027	0.238	2.166	0.024
Total	1.000	2.828	0.013	1.000	2.736	0.013
Number of observations	19,131	3,248		18,232	2,997	
1998						
Larger than average	0.151	3.312	0.018	0.128	3.309	0.021
Average	0.621	2.820	0.010	0.610	2.771	0.011
Smaller than average	0.181	2.276	0.020	0.210	2.271	0.020
Very small	0.048	1.928	0.049	0.053	1.828	0.047
Total	1.000	2.810	0.010	1.000	2.753	0.010
Number of observations	17,188	4,608		15,721	3,971	
2005						
Very large	0.043	3.389	0.046	0.039	3.149	0.050
Larger than average	0.192	3.078	0.015	0.186	2.975	0.018
Average	0.553	2.903	0.009	0.535	2.811	0.008
Smaller than average	0.151	2.300	0.020	0.164	2.289	0.020
Very small	0.060	1.930	0.032	0.075	1.938	0.031
Total	1.000	2.844	0.008	1.000	2.737	0.008
Number of observations	16,001	7,090		14,745	6,211	

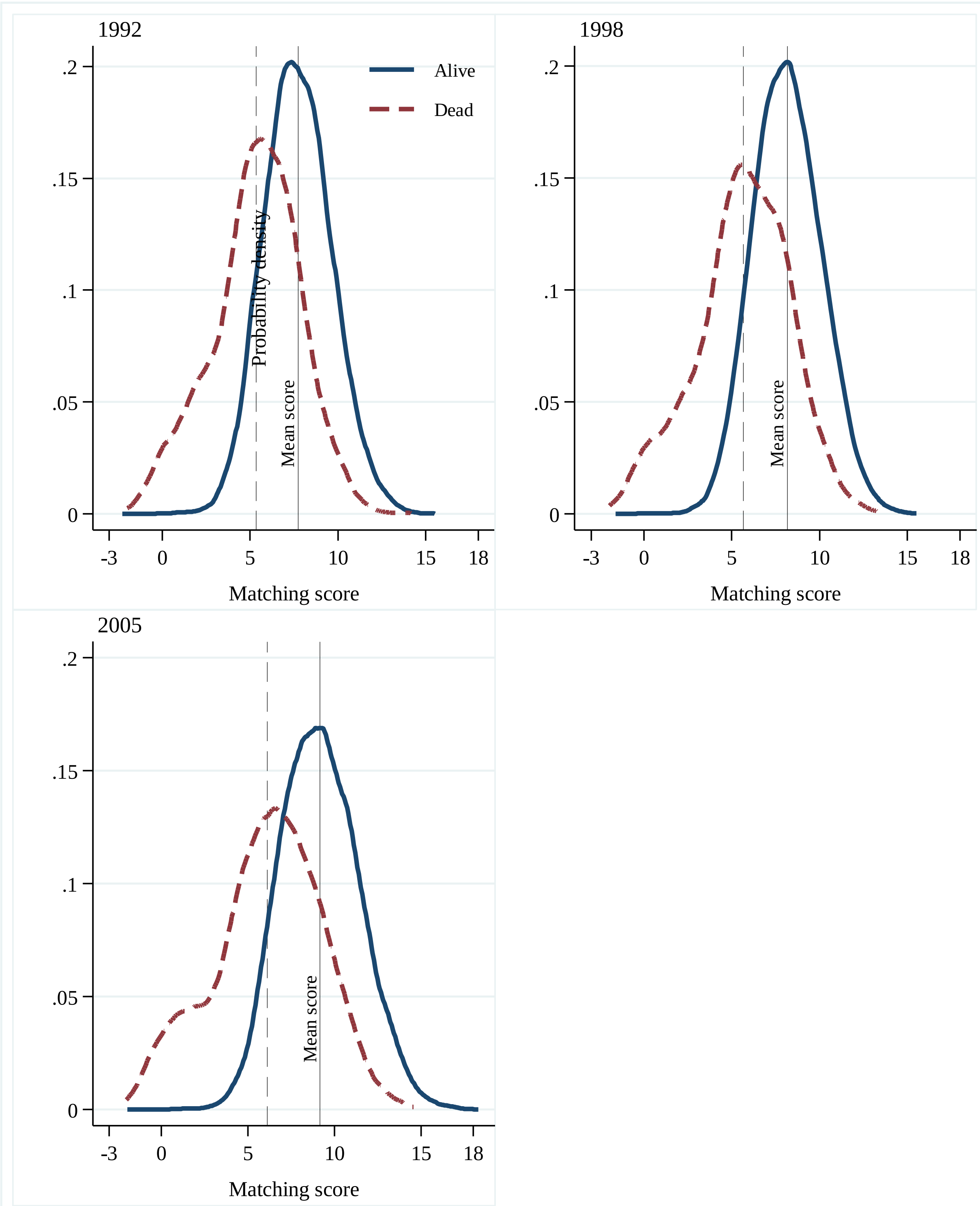
Source: Authors' analysis based on data from India's National Family Health Survey.

Table A2: Difference between actual and simulated height-for-age z-scores by gender for three rounds of India's National Family Health Survey, children ages 3–36 months

	Boys				*Girls*			
	Imputed HAZ	Mean HAZ Actual → Simulated	Δ	*Std. Error*	Imputed HAZ	Mean HAZ Actual → Simulated	Δ	*Std. Error*
	Sample of children with weight at birth assessed by their mothers							
1992	-2.106	-2.015 → -2.027	0.012	0.023	-2.145	-1.962 → -1.986	0.024	0.024
1998	-2.071	-1.877 → -1.893	0.016	0.020	-2.063	-1.955 → -1.964	0.009	0.022
2005	-1.630	-1.622 → -1.623	0.001	0.020	-1.725	-1.614 → -1.622	0.007	0.022
	Sample of children with measured weight at birth							
1992	-1.473	-1.360 → -1.368	0.008	0.050	-1.381	-1.374 → -1.374	0.000	0.051
1998	-1.764	-1.373 → -1.389	0.015	0.034	-1.743	-1.348 → -1.360	0.011	0.038
2005	-1.532	-1.246 → -1.257	0.011	0.029	-1.551	-1.239 → -1.249	0.010	0.031

Note: The standard errors for the differences are not adjusted for the fact that propensity score is estimated.
Source: Authors' analysis based on data from India's National Family Health Survey.

Figure A1: Kernel density estimates of distributions of the matching scores for samples of living and dead children for three rounds of India's National Family Health Survey



Source: Authors' analysis based on data from India's National Family Health Survey.